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TESTING THE MARSHALL-LERNER CONDITION IN KENYA

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ABSTRACT

In this paper we examine the Marshall-Lerner (ML) condition for the Kenyan economy. In particular, we use quarterly data on the log of real exchange rates, export-import ratio and relative (US) income for the time period 1996q1 – 2011q4, and employ techniques based on the concept of long memory or long-range dependence. Specifically, we use fractional integration and cointegration methods, which are more general than standard approaches based exclusively on integer degrees of differentiation. The results indicate that there exists a well-defined cointegrating relationship linking the balance of payments to the real exchange rate and relative income, and that the ML condition is satisfied in the long run although the convergence process is relatively slow. They also imply that a moderate depreciation of the Kenyan shilling may have a stabilizing influence on the balance of payments through the current account without the need for high interest rates.

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1.   Introduction

A lot of the literature on the balance of trade is based on the so-called "elasticity approach", namely on testing the extent to which trade flows are responsive to relative price changes, more specifically whether a devaluation improves the trade balance, which implies that the well-known Marshall-Lerner (ML) condition holds. The seminal empirical paper by Houthakker and Magee (1969) found inconclusive evidence. Several subsequent studies using least-squares methods to estimate price elasticities in import and export equations also produced mixed results (see, e.g., Khan 1974, Goldstein and Khan 1985, Wilson and Takacs 1979, Warner and Kreinin 1983, Bahmani-Oskooee 1986, Krugman and Baldwin 1987). More recently, the evidence obtained with more advanced econometric techniques taking into account non-stationarities in the data has been more supportive of the ML condition (see, e.g., Bahmani-Oskooee 1998, Bahmani-Oskooee and Niroomand 1998, Caporale and Chui 1999, Boyd, Caporale and Smith, 2001). Also, increasingly empirical investigations have been based on a reduced-form equation for the balance of trade, a method which allows to test directly for the response of trade flows to relative price movements using the real exchange rate (as opposed to the terms of trade) (see, e.g., Rose 1991, and Lee and Chinn 1998).

It is normally thought that a nominal devaluation or depreciation can only reduce trade imbalances if it translates into a real one and if trade flows respond to relative prices in a significant and predictable manner (Reinhart, 1995). A depreciation (or devaluation) of the domestic currency may stimulate economic activity through an initial increase in the price of foreign goods relative to home goods: by increasing the global competitiveness of domestic industries it diverts spending from the former to the latter (Kandil and Mirazaie, 2005).

Dornbusch (1988) shows that the effectiveness of a depreciation in improving the balance of payments depends on redirecting demand in the right direction and by the correct amount and also on the capacity of the domestic economy to meet the additional demand through increased supply. Bird (2001) argues that there is no mechanism for keeping the real exchange rate at an equilibrium level if inflation is rising quickly or for changing equilibrium rates in the case of permanent real shocks. In his opinion, this is the reason why many developing countries have chosen flexible exchange rates, although this is not an ideal solution since demand and supply elasticities may be relatively low: even when they satisfy the Marshall-Lerner conditions, their response to exchange rate changes may not be as big as in developed economies. Moreover, with thin foreign exchange markets floating exchange rates may be unstable and vulnerable to speculative attacks as the Kenyan exchange rate crisis of 2011 illustrated (Mudida, 2012): if the exchange rate is driven down sufficiently, it can generate additional inflation which may offset the extra competitiveness associated with a depreciation.

A related issue is whether there exist J-curve effects, i.e. whether following a currency depreciation or devaluation the balance of trade will worsen in the short run, but then, as elasticities increase over time, it will begin to improve (Kulkarni and Clarke, 2009). Case studies on several African countries such as Zambia and Nigeria seem to support empirically the existence of a J-curve. However, no studies have been carried out yet to test the Marshall-Lerner condition in Kenya, which represents an interesting case since its exports are relatively more diversified than those of the sub-Saharan African economies. The present paper is the first to provide evidence for this country; moreover, it uses advanced techniques in time series analysis, based on the concepts of fractional integration and cointegration, which are more general than the

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standard methods based exclusively on integer degrees of differentiation and have not been previously used to analyse the Marshall-Lerner condition in an African context.

The layout of the paper is as follows. Section 2 discusses the importance of the Marshall-Lerner condition in the Kenyan case. Section 3 briefly describes the theoretical framework, whilst Section 4 presents the econometric analysis. Finally, Section 5 summarises the main findings and offers some concluding remarks.

2.   The Marshall-Lerner condition and the Kenyan economy

The International Monetary Fund (IMF) classifies Kenya as having operated an independent float between 1992 and 1997 and a managed float since 1998. Prior to that, the Kenyan shilling was pegged first to the British pound, then to the US dollar, and finally to the IMF's Special Drawing Rights (SDRs) before a crawling peg based on a trade-weighted basket was introduced. The Marshall-Lerner condition should therefore be analysed in Kenya in the context of the current exchange rate system, which is a managed float system, and indeed the data set used in this study covers the floating period. Consequently, we consider a depreciation rather than a devaluation of the Kenyan shilling since this is what is relevant for the period under investigation. The existing empirical evidence on the operation of Kenya's managed float system suggests that at times of relative tranquillity in foreign exchange markets the Central Bank of Kenya can smooth out exchange rate volatility with relatively modest interventions; by contrast, more active policies are required in the presence of more volatile exchange rates (O'Connell et. al, 2010).

Testing the Marshall-Lerner condition is particularly important in the Kenyan case because, as in many other developing countries, the current account of the Kenyan balance of payments is persistently in deficit. The issue of whether a depreciation of the

exchange rate can reduce this deficit therefore becomes critical. A large share of Kenyan exports is represented by agricultural products with low price elasticities of demand. However, horticultural products and manufactured goods are also exported. Indeed, Kenya's largest trading partner is Uganda which imports primarily manufactured goods from Kenya. On the other hand, Kenyan imports are primarily made up of agricultural machinery, petroleum and manufactured goods which one would expect to have a low price elasticity of demand for imports owing to their critical role in the development process. This raises the interesting question of the size of real exchange depreciation required to eliminate Kenya's balance of payments current account deficit.

The broader issue related to the Marshall-Lerner condition in Kenya is that of real exchange targets for the Central Bank of Kenya, namely is there an optimal real exchange rate that should be targeted? A real exchange rate appreciation redirects resources from the export-producing sector penalizing it and causing potentially severe welfare losses (Rodrik, 2008). Pollin and Heintz (2007) have advocated the adoption of a new monetary policy framework in Kenya in order to achieve a more sizeable depreciation of the shilling in real terms. They stress that the contribution of the export sector, which is favoured by a real exchange depreciation, is unique from a development perspective owing to the number and quality of jobs created and also to the productivity spillovers to other sectors of the economy. The challenge of an excessively weak Kenya shilling, however, was illustrated in 2011 when a vicious cycle was created between inflation and depreciation (Mudida, 2012). This paper therefore also aims to contribute to the debate on the target for the real exchange rate of the Kenya shilling, which is a particularly interesting one because the primary task of the Central Bank of Kenya is price stability, at present being pursued through inflation targeting, with expected

inflation as the nominal anchor. It is well known that inflation targeting may be counterproductive in the presence of supply-side shocks, which are prevalent in the Kenyan economy (Adam et. al, 2010). Therefore analysing the Marshall-Lerner condition in Kenya is also important in view of the concerns facing the Kenyan monetary authorities.

3.   Theoretical Framework

The balance of trade can be expressed as the ratio of nominal exports to nominal imports, B, which is equal to the ratio of the volume of exports, X, multiplied by domestic prices, P, to the volume of imports M, multiplied by foreign prices, P^*, and the nominal spot exchange rate S:

$$B_t = \frac{P_t\, X_t}{P_t^*\, S_t\, M_t},$$

or using lower case letters for logarithms:

$$b_t = x_t - m_t - \left(s_t - p_t + p_t^*\right) = x_t - m_t - e_t, \qquad (1)$$

where $e_t = s_t - p_t + p_t^*$ is the real exchange rate. Long-run import and export demand are given by:

$$x_t = \alpha_x + \beta^* y_t^* + \eta_x e_t + \gamma_x t, \qquad (2)$$

$$m_t = \alpha_m + \beta y_t - \eta_m e_t + \gamma_m t. \qquad (3)$$

where y_t and y_t* stand for domestic and foreign real income respectively, the trends capture terms of trade effects, and η_x and η_m represent the export and income elasticities respectively.

The long-run balance of trade is

$$b_t = (\alpha_x - \alpha_y) + \beta^* y_t^* - \beta y_t + (\eta_x + \eta_m - 1)e_t + (\gamma_x - \gamma_m)t \qquad (4)$$

The coefficient on e_t gives the familiar Marshall-Lerner condition for a devaluation (increase in e_t) to improve the balance of payments (i.e., this coefficient needs to be statistically significant and positive for the ML condition to be satisfied, which means that the sum of the demand elasticity for imports and the foreign demand elasticity for the nation's export exceeds unity). Solvency requires $b_t = 0$ in the long run, whilst Purchasing Power Parity (PPP) requires $e_t = e$, for all t.

The long-run relationship (4) can be written as

$$b_t = \alpha + \beta^* y_t^* - \beta y_t + \eta e_t + \gamma t, \qquad (5)$$

where $\alpha = (\alpha_x - \alpha_m)$; $\eta = (\eta_x - \eta_m - 1)$, and $\gamma = (\gamma_x - \gamma_m)$, and the deviations from the long-run equilibrium can be defined as

$$z_t = \alpha + \beta^* y_t^* - \beta y_t + \eta e_t + \gamma t - b_t. \qquad (6)$$

4.   Econometric Analysis

We use quarterly data on the log of real exchange rates, export-import ratio and relative (US) income for the time period 1996q1 – 2011q4. All series were obtained from the Central Bank of Kenya. The base period for the real effective exchange rate is January 2003. The real effective exchange rate is constructed using a basket including the eight countries which are Kenya's most important trading partners, and is defined in such a way that an increase represents a depreciation, i.e. a direct quote is used as common in developing countries.

Figure 1 displays the three series (in logs) in both levels and first differences. The export/import ratio and the real exchange rate decline over the sample period, whilst relative income increases. The first differenced series data show that seasonality is an important feature of these data, especially for the export/import ratio and relative income.

[Insert Figures 1 – 3 about here]

Figure 2 displays the first thirty sample autocorrelations for the original series and their first differences. The slow decay for the series in levels indicates that they may be nonstationary, while the sample autocorrelations for the first differences suggest once more the presence of seasonality, especially in the case of relative income. Finally, Figure 3 displays the periodograms. For the series in levels the highest value corresponds to the smallest frequency, which indicates that they may require differencing. However, the periodogram of the first differenced export/import ratio series has a value close to zero at the smallest frequency, suggesting that this series may now be overdifferenced.

As a first step we check the order of integration of the three series by means of standard methods such as ADF (Dickey and Fuller, 1979), Phillips and Perrron (PP, 1988), Kwiatkowski et al. (KPSS, 1992), Elliot et al. (ERS, 1996) and Ng and Perron (NG, 2001) tests. The results (not reported) are conclusively in favour of unit roots for the real exchange rate and relative income, whilst mixed evidence is found in the case of the export/import ratio. However, they should be taken with caution, since the above methods have limitations such as very low power if the true Data Generating Process (DGP) is fractionally integrated (see, e.g., Diebold and Rudebusch, 1991; Hassler and Wolters, 1994; Lee and Schmidt, 1996 among others). Thus, in what follows we consider models that allow for both integer and fractional orders of differentiation.

We estimate d (the differencing parameter) using the Whittle function in the frequency domain (Dahlhaus, 1989) and also employ a testing procedure developed by Robinson (1994) which has been shown to be the most efficient in the context of I(d) models. The latter method is parametric, so a parametric model for the disturbances

term has to be specified. A semiparametric method, also based on the Whittle function (Robinson, 1995, Abadir et al., 2007) will also be employed.

We report in Table 1 the estimated values of d in a model given by

$$y_t = \alpha + \beta t + x_t; \quad (1-L)^d x_t = u_t, \quad (7)$$

where y_t is the observed (univariate) time series; α and β are the coefficients on the intercept and a linear trend respectively, and x_t is assumed to be an I(d) process. Thus, u_t is I(0) and given the parametric nature of this method its functional form must be specified. We assume that u_t is a white noise, autocorrelated and seasonally autoregressive respectively. In the case of autocorrelated errors, we use the exponential model of Bloomfield (1973). This is a non-parametric approach for modelling u_t that produces autocorrelations decaying exponentially as in the AR(MA) case. The model is implicitly determined by its spectral density function, which is given by:

$$f(\lambda;\tau) = \frac{\sigma^2}{2\pi} \exp\left(2 \sum_{r=1}^{m} \tau_r \cos(\lambda r) \right), \quad (8)$$

where σ^2 is the variance of the error term, and m is the number of parameters required to describe the short-run dynamics of the series. Bloomfield (1973) showed that the logarithm of an estimated spectral density function is often a fairly well-behaved function and can thus be approximated by a truncated Fourier series; in particular, the spectral density of an ARMA process can be well approximated by (8). Moreover, this model is stationary across all values of τ, and works extremely well in the context of Robinson's (1994) tests (Gil-Alana, 2004).

[Insert Table 1 about here]

Table 1 displays the estimates of d (along with the 95% confidence bands corresponding to the non-rejection values of d using Robinson's (1994) tests) for the three types of disturbances (white noise, Bloomfield, and seasonal AR) and for the three

standard cases of: i) no regressors (i.e., α = β = 0 a priori in (7)), an intercept (α unknown and β = 0 a priori), and an intercept with a linear time trend (i.e., α and β unknown). The t-values (not reported) for the deterministic terms indicate that the time trend is required in all cases. The upper part of the table refers to the case of white noise disturbances. Focusing on the case of a linear trend, we see that the estimated value of d for the log(export/import) ratio is 0.373, and the confidence interval excludes the cases of stationarity I(0) (d = 0) and nonstationary unit roots (d =1). For the real exchange rate, the estimated d is 0.888 and the unit root null hypothesis cannot be rejected at the 5% level. Finally, for relative income, the estimated value of d is 0.664 and the I(1) case is rejected in favour of mean reversion (d < 1). The results based on the assumption of autocorrelation as in the model of Bloomfield (1973) (with m = 1)[1] are displayed in Table 1(ii). The estimated values of d (for the case of a linear trend) are 0.574, 0.728 and 0.865 respectively for the export/import ratio, real exchange rates and relative income, and the unit root null cannot be rejected in the last two cases. Finally, when imposing seasonal (quarterly) autoregressions, these values are 0.376, 0.899 and 0.963 and similarly to the previous case the I(1) hypothesis cannot be rejected for real exchange rates and relative income, while it is rejected in favour of mean reversion for the export/import ratio.

The above results are corroborated by those based on the semiparametric method of Robinson (1995). This is essentially a local 'Whittle estimator' in the frequency domain, which uses a band of frequencies degenerating to zero. The estimator is implicitly defined by:

$$\hat{d} = \arg\min_{d} \left(\log \overline{C(d)} - 2d \, \frac{1}{m} \sum_{s=1}^{m} \log \lambda_s \right), \qquad (9)$$

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[1] Other values of m produced essentially the same results.

$$\overline{C(d)} \;=\; \frac{1}{m} \sum_{s=1}^{m} I(\lambda_s)\, \lambda_s^{2d}, \qquad \lambda_s = \frac{2\pi s}{T}, \qquad \frac{1}{m} + \frac{m}{T} \to 0,$$

where $I(\lambda_s)$ is the periodogram of the raw time series, and $d \in$ (-0.5, 0.5). Under finiteness of the fourth moment and other mild conditions, Robinson (1995) proved that:

$$\sqrt{m}\,(\hat{d} - d^*) \to_{dtb} N(0, 1/4) \quad as\ T \to \infty,$$

where d^* is the true value of d. This estimator is robust to a certain degree of conditional heteroscedasticity (Robinson and Henry 1999) and is more efficient than other semi-parametric competitors.[2]

[Insert Figure 4 about here]

Figure 4 display the estimates of d in (9) along with the 95% confidence interval corresponding to the I(0) and I(1) cases. The horizontal axis reports the bandwidth parameter while the vertical one the estimates of d. For the export/import ratio some of the estimates are within the I(1) interval but most are below it although above the I(0) one. For the real exchange rate, most of the estimates of d are within the I(1) interval. Finally, for relative income, they are above the I(1) interval if the bandwidth parameter is low, whilst are within it if it is large. This may be a consequence of the strong seasonal pattern observed in this series.[3] It is also consistent with the results reported for the parametric case above where the unit root cannot be rejected in case of the exchange rate (and in some cases for relative income) and is rejected in favour of mean reversion for the export/import ratio.

Considering again the results presented in Table 1, we next select the best model specification for each series. We conducted several diagnostic tests on the residuals of the estimated models, and, in particular, we used Box-Pierce and Ljung-Box-Pierce

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[2] This method has been further refined by Velasco (1999), Velasco and Robinson (2000), Phillips and Shimotsu (2004, 2005), Abadir et al. (2007) and others. When using these approaches the results were practically identical to those reported in the paper.
[3] The bandwidth determines the trade-off between the bias and the variance in the estimation of d.

statistics (Box and Pierce, 1970; Ljung and Box, 1978) to test for no serial correlation, as well as LR tests and other likelihood criteria.[4] The selected models for each variable are the following:

$$y_t = -0.4073 - 0.0081t + x_t, \quad (1-L)^{0.573} x_t = u_t,$$
$$\quad\;\; (-5.09) \quad\;\; (-3.00)$$
$$u_t \approx \text{Bloomfield } (\tau = -0.291)$$

for the export/import ratio. (t-values in parenthesis).

For the real exchange rate, the selected model is

$$y_t = 4.6981 - 0.0064t + x_t, \quad (1-L)^{0.728} x_t = u_t,$$
$$\quad\;\; (135.12) \quad\;\; (-3.71)$$
$$u_t \approx \text{Bloomfield } (\tau = 0.208)$$

Finally, for relative income,

$$y_t = -3.8257 - 0.0106t + x_t, \quad (1-L)^{0.963} x_t = u_t,$$
$$\quad\;\; (-204.33) \quad\;\; (-5.19)$$
$$u_t = 0.893 u_{t-4} + \varepsilon_t .$$

The fact that the confidence intervals for the fractional differencing parameters in the selected models overlap for the three series[5] implies that the null of equal orders of integration cannot be rejected. This is important since it makes it legitimate to run an OLS regression with the three variables to check if the estimated errors are I(0) or at least mean-reverting with a smaller order of integration than the three parent series.[6]

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[4] Specifically, the AIC and the SIC. Note, however, that these criteria might not necessarily be the best ones in applications involving fractional differences, as they focus on the short-term forecasting ability of the fitted model and may not give sufficient attention to their long-run properties (see, e.g. Hosking, 1981, 1984).

[5] These intervals are (0.267, 0.975) for the export/import ratio, (0.441, 1.124) for the real exchange rate, and (0.763, 1.247) for relative income (see Table 1).

[6] We also compute an adaptation of the Robinson and Yajima (2002) statistic $\hat{T}_{xy}$ for log-periodogram estimation in pairwise comparisons of the three series; the results support the hypothesis of homogeneity in the orders of integration.

We follow a two-step procedure, similar to that of Engle and Granger (1987), but specifically designed to allow for fractional integration. In the first step, we compute the following regression,

$$y_t = \alpha + \beta_1 z_{1t} + \beta_2 z_{2t} + x_t, \tag{10}$$

where y_t stands for the balance of trade, z_{1t} for the real exchange rate and z_{2t} for relative income. Then, in the second step, we estimate the order of integration of the residuals from (10) using the methods employed above for the univariate analysis.

Performing the OLS regression in (10) we obtain

$$y_t = -3.4087 + 0.0942 z_{1t} - 0.6643 z_{2t} + \hat{x}_t .$$
$$\quad\; (-6.84) \qquad (1.77) \qquad\quad (-4.89) \tag{11}$$

The estimated residuals are plotted in Figure 4. The positive (and statistically significant) coefficient on the real exchange rate indicates that the ML condition is satisfied in Kenya.

[Insert Figure 4 and Table 2 about here]

Table 2 displays the estimated values of d for the OLS residuals. We consider again the three standard cases of no regressors, an intercept and a linear trend, for white noise, Bloomfield, and seasonal AR disturbances. The estimates are all in the range (0.23, 0.25) being substantially smaller than those for the individual series and thus supporting the existence of mean reversion in the long-run equilibrium relationship. If we focus now on the confidence intervals we see that the null hypothesis of I(0) errors is rejected for the cases of white noise and seasonal AR disturbances in favour of positive orders of integration, while this hypothesis cannot be rejected with autocorrelated (Bloomfield) disturbances.

The most adequate specification for the estimated residuals in the cointegrating regression (10) appears in bold in Table 2. This model includes an intercept and

seasonal AR disturbances. We also perform the Hausman-type test of no cointegration against the alternative of fractional cointegration proposed by Marinucci and Robinson (2001); the results (not reported) strongly reject the null for different bandwidth parameters given further support to the hypothesis of cointegration among the variables examined.

5.   Conclusions and Policy Recommendations

Our findings support the existence of a well-defined cointegrating relationship between the balance of payments, the real exchange rate and relative income and indicate that the Marshall-Lerner condition holds in Kenya. Our analysis is based on fractional integration and cointegration methods, which are more general than the standard methods allowing only for integer degrees of differentiation. Studies using the latter to test the Marshall- Lerner condition in many African countries are either inconclusive or tend to suggest that it does not hold in the short run although it may hold in the long run. The evidence (based on more general methods) that it holds in Kenya has important policy implications for this country. It implies that the exchange rate is an important tool for attempting to address persistent balance of payments current account deficits in Kenya and can therefore contribute to achieving an external balance.

The fact that the Marshall-Lerner condition holds means that a depreciation of the exchange rate leads to a reduction in import expenditure and an increase in export sales. This reflects an important transition made in Kenya in terms of the composition of exports: from traditional agricultural exports exhibiting low export elasticities of demand to more diversified non-traditional exports such as horticulture and manufactured goods that exhibit a higher elasticity of demand. Our results indicate that a depreciation in the Kenya shilling can therefore have potentially beneficial effects on

Kenya's current account deficit. These, however, have to be weighed against the higher inflation rate associated with such an exchange rate movement. Inflationary effects were evident during the 2011 foreign exchange crisis in Kenya when a depreciation of 30% of the Kenya shilling against the US dollar led to month-on-month inflation of 19% by the end of 2011 (Mudida, 2012).

Given that at present the primary objective of the Central Bank of Kenya is price stability, the focus recently has been on maintaining high interest rates so as to reduce the inflation rate and also to avoid a significant depreciation of the Kenya shilling. This tight monetary policy stance is thought to reduce inflationary pressure and also to promote portfolio investment inflows into Kenya, thus improving the capital account. High interest rates tend to have a detrimental effect on economic growth. Our findings, however, suggest that a moderate depreciation of the Kenyan shilling may in fact have a stabilizing influence on the balance of payments through the current account without the need for high interest rates. Thus a less contractionary monetary policy by the Central Bank of Kenya could in fact be combined with an appropriate exchange rate policy to achieve more effectively the objectives of internal and external balance in Kenya. This would be a better option than the current high interest rate policy being pursued by the Central Bank that achieves external balance but only at the high cost of stifling economic growth.

Other recently developed bivariate or multivariate fractional cointegration testing methods (e.g. Johansen, 2010; Nielsen, 2010; Nielsen and Frederiksen, 2011) could also be applied. This could be very useful to investigate possible J-curve effects in the context of a much richer structure including other underlying dynamics and short-run components. Structural breaks and non-linearities could also be examined in the context of fractional integration. These issues will be investigated in future papers.

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**Figure 1: Original series and first differences**
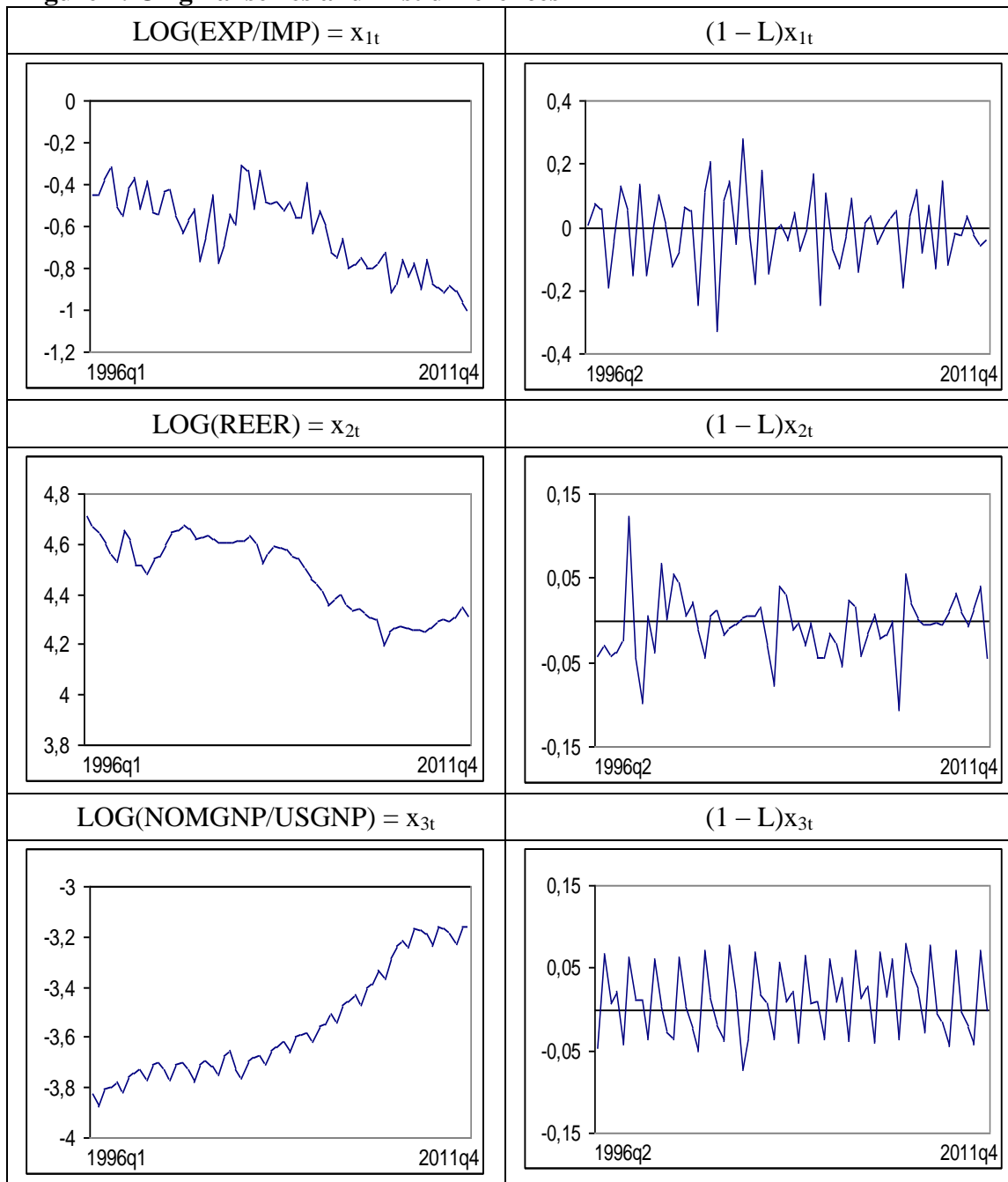
EXP/IMP = Export/Import ratio; REER = Real Effective Exchange Rate; NOMGNP = Nominal GNP and USGNP= US GNP.

Figure 2: Correlograms of the original series and first differences

LOG(EXP/IMP) = x_{1t}	$(1-L)x_{1t}$
LOG(REER) = x_{2t}	$(1-L)x_{2t}$
LOG(NOMGNP/USGNP) = x_{3t}	$(1-L)x_{3t}$
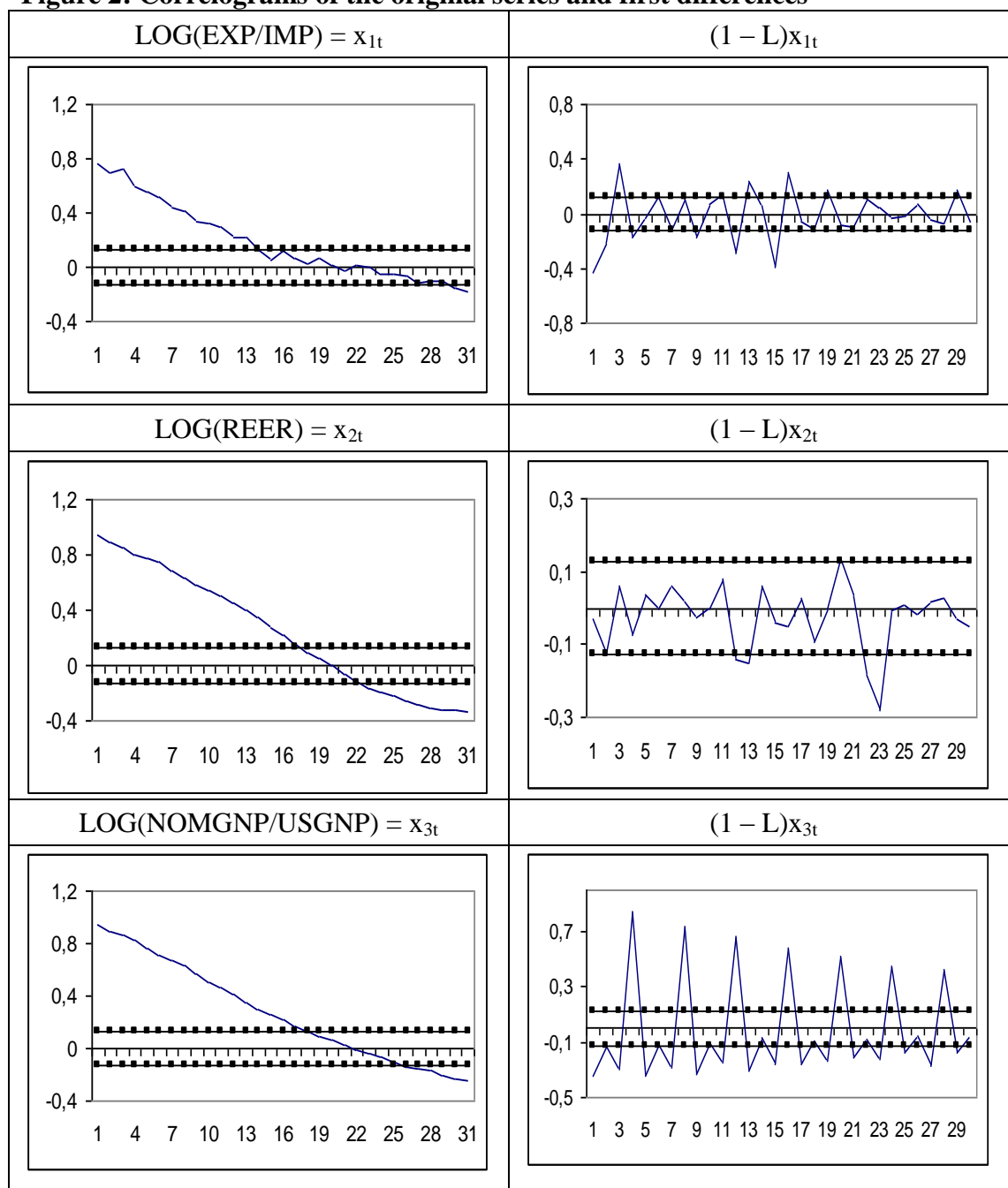
The thick lines give the 95% confidence band for the null hypothesis of no autocorrelation.

**Figure 3: Periodograms of the original series and first differences**

LOG(EXP/IMP) = x_{1t}	$(1-L)x_{1t}$

LOG(REER) = x_{2t}	$(1-L)x_{2t}$

LOG(NOMGNP/USGNP) = x_{3t}	$(1-L)x_{3t}$
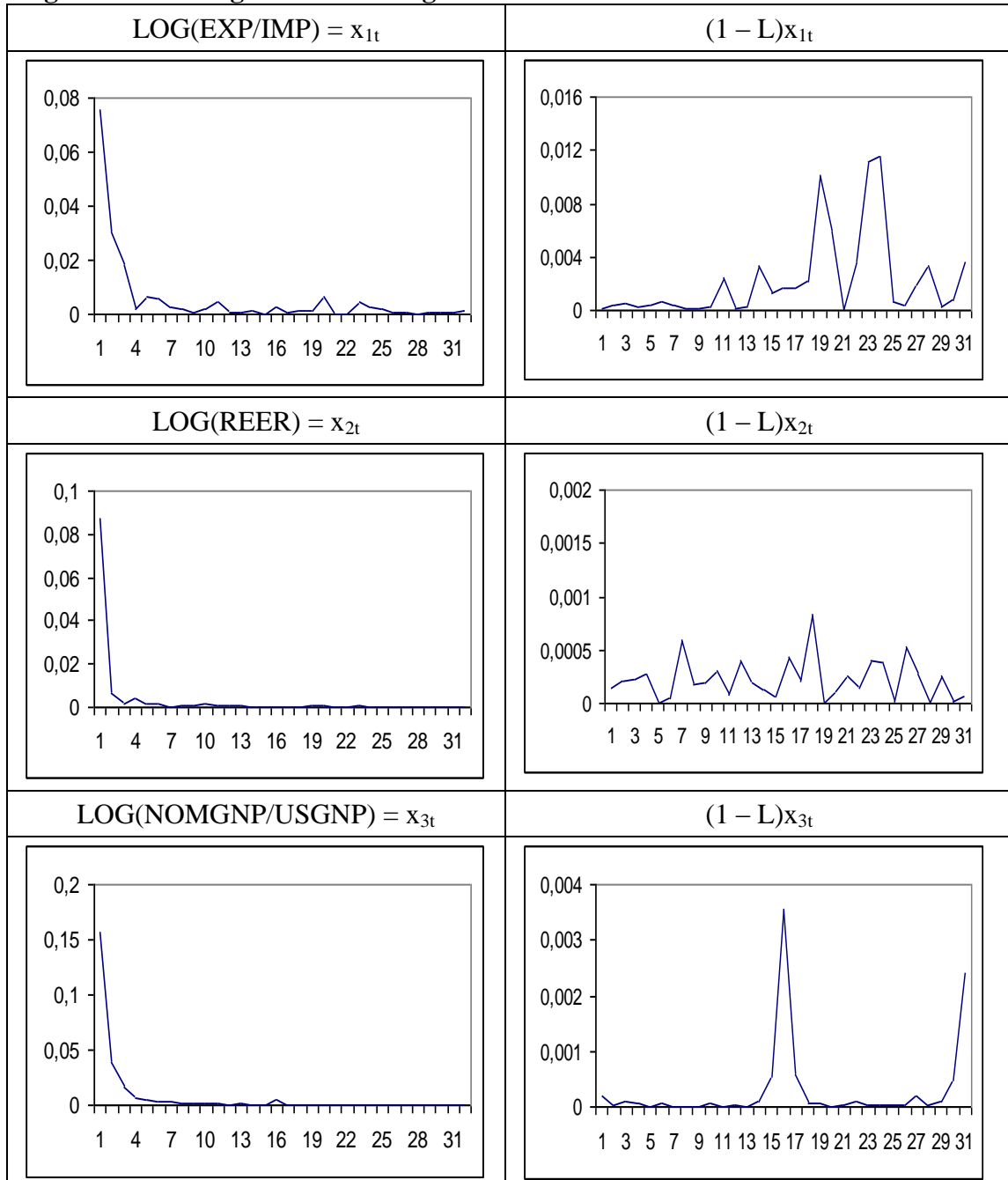
The horizontal axis refers to the discrete Fourier frequencies $\lambda_j = 2\pi j/T$, $j = 1, \ldots, T/2$.

Table 1: Estimates of d and 95% confidence bands for the three individual series

	No regressors	An intercept	A linear time trend
i) White noise disturbances			
LOG(EXP/IMP)	0.511 (0.372, 0.714)	0.493 (0.413, 0.605)	**0.373 (0.258, 0.536)**
LOG(REER)	0.934 (0.787, 1.148)	0.883 (0.742, 1.129)	**0.888 (0.740, 1.129)**
LOG(NOM/USGNP)	0.949 (0.792, 1.174)	0.755 (0.691, 0.856)	**0.664 (0.569, 0.802)**
ii) Bloomfield-type disturbances			
LOG(EXP/IMP)	0.728 (0.394, 1.144)	0.643 (0.422, 0.971)	**0.574 (0.267, 0.975)**
LOG(REER)	0.842 (0.552, 1.234)	0.803 (0.637, 1.117)	**0.728 (0.441, 1.124)**
LOG(NOM/USGNP)	0.824 (0.543, 1.173)	0.892 (0.741, 1.074)	**0.865 (0.673, 1.093)**
iii) Seasonal (quarterly) AR disturbances			
LOG(EXP/IMP)	0.495 (0.321, 0.734)	0.495 (0.361, 0.626)	**0.376 (0.223, 0.569)**
LOG(REER)	0.857 (0.546, 1.155)	0.896 (0.745, 1.134)	**0.899 (0.723, 1.137)**
LOG(NOM/USGNP)	0.885 (0.574, 1.173)	0.969 (0.802, 1.262)	**0.963 (0.763, 1.247)**

Figure 4: Estimates of d and 95% confidence bands for the three individual series

i) LOG(EXP/IMP)

ii) LOG(REEF)

iii) Log(NOM/USGNP)
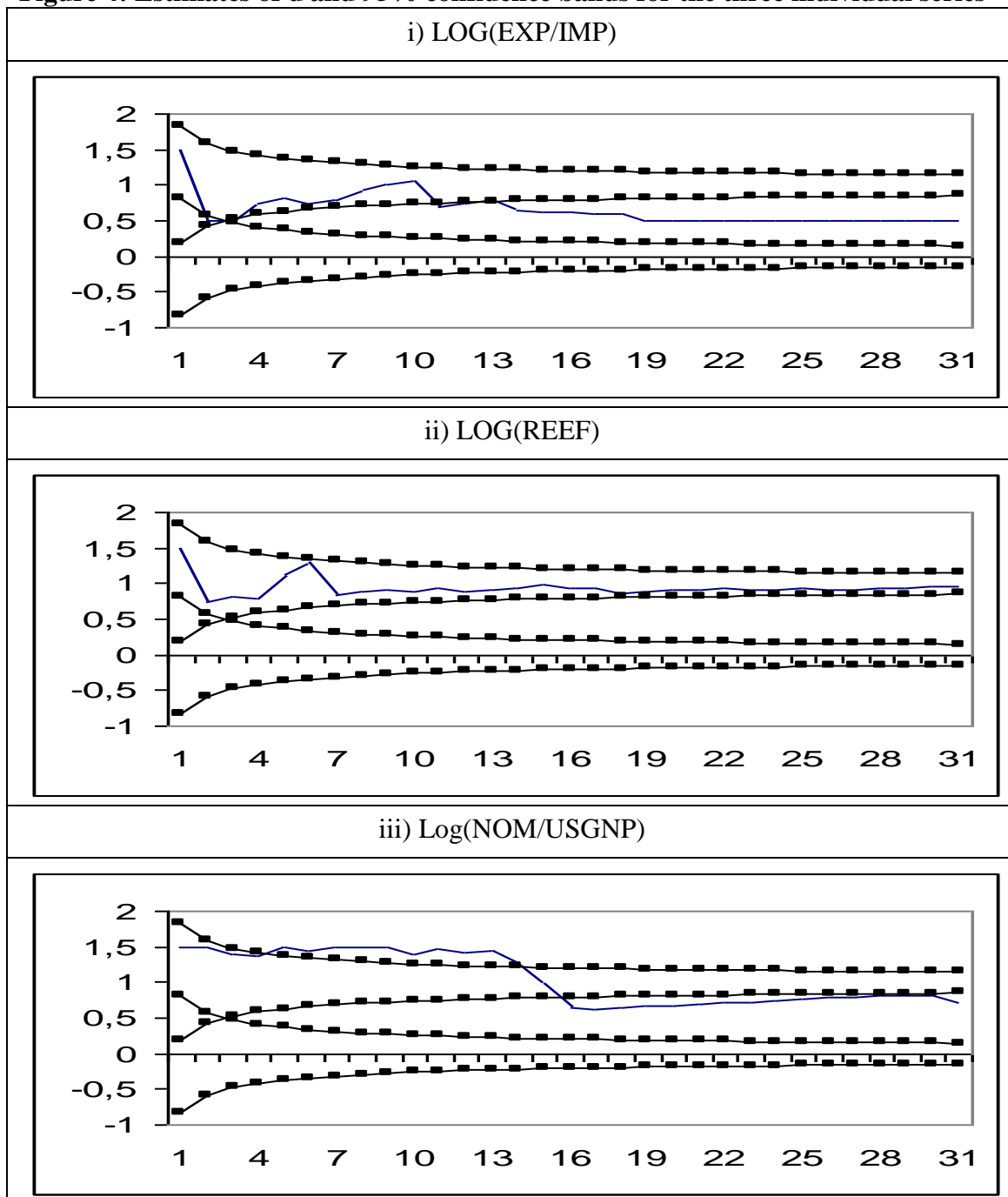
The horizontal axis concerns the bandwidth parameter while the vertical one refers to the estimated value of d. The thick lines give the 95% confidence bands for the I(0) and I(1) hypotheses.

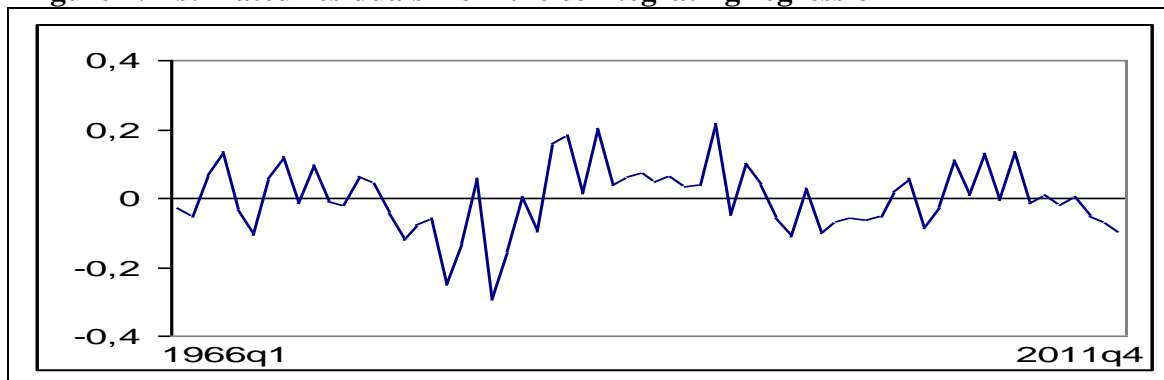**Figure 4: Estimated residuals from the cointegrating regression**

Table 2: Estimates of d and 95% confidence bands for the three individual series

i) White noise disturbances			
	No regressors	An intercept	A linear time trend
White noise	0.239 (0.089, 0.435)	0.239 (0.089, 0.434)	0.241 (0.091, 0.436)
Bloomfield	0.255 (-0.046, 0.575)	0.258 (-0.050, 0.579)	0.259 (-0.048, 0.579)
Seasonal AR	0.244 (0.072, 0.453)	0.244 (0.071, 0.452)	0.246 (0.072, 0.455)